SOCIOLOGICAL METHODOLOGY

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# ESTIMATING NET INTERRACIAL MOBILITY IN THE UNITED STATES: A RESIDUAL METHODS APPROACH

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*This paper presents a residual methods approach to identifying social mobility across race/ethnic categories. In traditional demographic accounting models, population growth is limited to changes in natural increase and migration. Other sources of population change are absorbed by the model residual and can be estimated only indirectly. While these residual estimates have been used to illuminate a number of elusive demographic processes, there has been little effort to incorporate shifts in racial identification into formal accounts of population change. In light of growing evidence that a number of Americans view race/ethnic identities as a personal choice, not as a fixed characteristic, mobility across racial categories may play important roles in the growth of race/ethnic subpopulations and changes to the composition of the United States. To examine this potential, we derive a reduced-form population balancing equation that treats fertility and international migration as given and estimates survival from period life table data. After subtracting out national increase and net international migration and adjusting for changes in racial measurement*

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*and census coverage, we argue that the remaining error of closure provides a reasonable estimate of net interracial mobility among the native born. Using recent U.S. Census and ACS microdata, we illustrate the impact that identity shifts may have had on the growth of race/ethnic subpopulations in the past quarter century. Findings suggest a small drift from the non-Hispanic white population into race/ethnic minority groups, though the pattern varies by age and between time periods.*

Population change is a natural preoccupation of sociological inquiry. The dynamics of natural increase and migration shape the size and structure of populations, and the interplay between these processes forms the core logic of demographic accounting. Armed with basic methods such as the population balancing and renewal equations, demographers routinely construct forecasts of future populations, provide estimates of migratory flows, evaluate the coverage and accuracy of census enumerations, and help shed light on population dynamics from the distant past using fragmentary data sources. In this paper, we extend these methods to measure net mobility across race and ethnic categories.

We begin by reviewing different ways in which populations can be defined, noting that race/ethnicity blurs the line between a fixed characteristic and a flexible identity. After outlining some conceptual and methodological challenges that race poses for traditional demographic accounting models, we devise a residual methods approach to estimating net flows between race/ethnic groups over time. We then illustrate the feasibility of the method using recent U.S. Census and American Community Survey data.

## 1. DEMOGRAPHIC ACCOUNTING AMONG DIFFERENT TYPES OF POPULATIONS

Though defined at particular points in time and space, populations are always in motion. Growth and decline largely hinge on the balance of two biological processes—fertility and mortality—though geographic mobility often plays an integral role. The interplay between these three processes governs the logic of demographic accounting, as shown in

$$\delta P = Births - Deaths + Immigrants - Emigrants, \tag{1}$$

which states that the total change in any population is simply the sum of natural increase (births minus deaths) and net migration (in-migrants minus out-migrants). This formal identity, known as the *population balancing equation*, is one of the oldest and most basic tools of demographic accounting.

The logic of the balancing equation easily extends to subpopulations defined by ascribed characteristics—those that remain fixed throughout the life course (sex, birthplace, etc.) or that change in a predictable pattern (e.g., age). Subpopulations defined by *achieved* characteristics present a more complex undertaking (Ryder 1964; Schnore 1961), since individuals can enter and exit these populations through changes in social status (e.g., from married to divorced). Both the number and timing of status transitions confound the simple logic of demographic accounting. A given individual can be counted in different populations (e.g., married versus unmarried) throughout the life course, and a given status (e.g., all unmarried persons) can apply to different types of persons (never married versus divorced). Still, with simplifying assumptions about life course changes, demographic accounting methods can be adapted to measure components of change within populations defined by achieved characteristics, such as occupation and religious affiliation (Duncan 1963; Fischer and Hout 2006, ch. 8).

While the accounting framework is robust, the methodology offers little insight on whether a particular population characteristic should be treated as ascribed or achieved. Race (or race/ethnicity),[1] for example, is widely included as an ascribed characteristic in demographic accounts. A popular demography textbook lists race, along with groups delineated by sex and nativity, as common examples of "populations defined by characteristics fixed at birth" (Preston et al. 2001:2). In keeping with this perspective, the Census Bureau routinely projects the future race/ethnic composition of the American population (U.S. Census Bureau 2008), and mainstream media outlets are often quick to relay these

[1] Unless otherwise noted, we use "race" and "race/ethnicity" interchangeably throughout the text. While race and ethnicity can be viewed as conceptually distinct (e.g., race is more often associated with physical appearance), the two concepts often serve the same function. Many white ethnic groups were considered distinct "races" or peoples at various points in time (Waters 1990), and many ethnic and national origin groups (Chinese, Korean, etc.) are defined as races on current census forms (see Appendix A). Likewise, Hispanics, the sole "ethnic" group defined by current guidelines (OMB 1997), are often treated as a distinct "race"(coequal with non-Hispanic whites, blacks, etc.) in published research and public consciousness (Logan 2004; Perez 2008).

projections as "official" forecasts of America's racial future some 20, 30, even 40 years out (Roberts 2008). These projections, and the coverage they receive, reinforce the idea that the racial identities recorded in the census and other data sources are exogenous to the individual and subject to uniform rules of inheritance. These assumptions raise a number of conceptual and practical concerns, however (Hirschman 2002).

## 2. THE CHALLENGE OF RACE/ETHNICITY FOR DEMOGRAPHIC ACCOUNTING

Concepts of race and ethnicity have experienced a dramatic shift away from once widely held beliefs that identities are solely a function of biological inheritance or ancestry. The critical backlash against ideological racism in the wake of World War II signaled the downfall of essentialist thinking about race (Fredrickson 2002), and in the decades that followed, there emerged a growing recognition of the cultural and political underpinnings of racial and ethnic identities (Hirschman 2004). Today, not only are race and ethnicity primarily viewed as social constructs, but there is also increasing evidence that race/ethnic identities can vary across generations, and possibly over the life course of individuals.

Though several factors contribute to the ambiguity of race/ethnic identities, none is more important than the blurring of group boundaries through population admixture. Ethnic blending is most often attributed to intermarriage and the gradual absorption of new immigrant groups. Accordingly, more nuanced projections of America's race/ethnic diversity include a *range* of estimates using alternate assumptions about future levels of immigration and intermarriage (Edmonston and Passel 1994; Edmonston et al. 2002; Gibson 1992). While these projections improve upon those released by the census, recent intermarriage may not fully capture the mixed ancestry of many Americans (Perez and Hirschman 2009), particularly those whose ethnic descent is shrouded in murky and often painful histories, such as the absorption of indigenous peoples and the prevalence of forced sexual unions between plantation-era property owners and slaves (Davis 1991; Spencer 2006).

Only fragments of Americans' race/ethnic complexity are reflected in the statistical record. Modern census-taking relies on "mail-in" questionnaires that are filled out by individuals, so the decision to report a particular identity, but not others, is largely a matter of personal choice. This pattern is particularly well documented for persons of mixed European ancestry, who comprise the majority of non-Hispanic whites in the United States (Alba 1990). Most whites oversimplify their ancestral roots on the census and social surveys, and the selective reporting of certain ancestries stems from a number of factors, including personal sentiments about favored relatives and holidays (Waters 1990), popular stereotypes about ethnic groups (Hout and Goldstein 1994; Perry 2002), and at times the mere listing of particular nationalities (e.g., Irish, Italian) on the census form, which some may interpret as suggestions for how to identify (Farley 1991).

With socioeconomic divisions between white ethnic groups having all but disappeared from contemporary American life (Alba 1990), the blurring of ethnic boundaries is not only reasonable, it is exactly what the popular image of a "melting pot" society would predict (Bonnett 1998; McDermott and Samson 2005). Racial identities, which are usually associated with physical appearance, are not as malleable as cultural or ancestral ties (Nagel 1996), though there are many similarities in the ways race and ethnicity have functioned throughout U.S. history (see footnote 1). While African ancestry is commonly associated with the one drop rule that makes "black" identity mandatory (Davis 1991; Hollinger 2005), many persons of mixed African and European descent identified as multiracial in the eighteenth and early nineteenth centuries (Williamson 1980), while others chose to "pass" into the white population to escape racial oppression (Davis 1991; Piper 1992). In an analysis of census race responses among children with white and Asian parents, Xie and Goyette (1997) find that the choice of race is largely *optional*, although language, immigrant generation, and surname all play a role in how biracial children are classified. Among Native Americans, population growth in recent censuses so vastly outpaced expectations that researchers could only attribute the excess growth to changes in racial self-classification among persons with partial indigenous ancestry (Eschbach 1993; Eschbach 1995; Harris 1994; Nagel 1996; Passel 1976, 1996; Passel and Berman 1986; Snipp 1997). Race and ethnicity also seem to be especially fluid among children transitioning to adulthood, as recent studies have shown that adolescents can change identities even within short time periods (Harris and Sim 2002; Perez 2008).

The ambiguities in race/ethnic reporting are compounded by frequent changes and inconsistencies in the system of race/ethnic measurement and classification (Ferrante and Brown 1998; Perez and Hirschman 2009; Perlmann and Waters 2002). Both the categories and methods by which race is measured vary between data sources and over time (Office of Management and Budget 1977, 1997; Schenker 2003). For example, Passel (1976) notes that the excess population growth among American Indians was first observed in the "mail-in" census of 1960—the first to heavily rely on self-classification, rather than enumerator observation, to identify the race of household members. The decision to allow respondents to report multiple races was also a major departure from earlier census forms (Farley 2002), as it increased the minimum number of categories from five in 1990 to more than 60 (due to combinations) in 2000. The Census Bureau implemented these changes in time for the 2000 Census, but other statistical agencies were given additional time to adopt the new rules (Office of Management and Budget 1997). Compliance remains uneven to this day.

As argued in earlier work (Perez and Hirschman 2009), the difficulties in defining and measuring race/ethnic identities cast suspicion on most, if not all, projections of the future race/ethnic diversity of the United States. The potential for mobility across categories challenges the conventional wisdom that treats race/ethnicity as an ascribed characteristic in population accounting models. At the same time, these challenges provide an opportunity to expand the demographic toolkit to incorporate identity shifts into formal accounts of population change. Doing so requires a new analytic starting point, however—one that explicitly recognizes that while race may be seen as an ascriptive attribute by the majority of the population, there is a minority, of unknown magnitude, for whom race is a social identity that can change over the life course.

With minor revisions, demographic accounting can accommodate both types of persons, as shown in

$$\delta P = \sum_{k=1}^{n} \delta P_k,$$

where $k$ denotes each race/ethnic group, and

$$\delta P_k = B_k - D_k + I_k - E_k + \varepsilon_k. \tag{2}$$

This simple decomposition of (1) acknowledges that most persons enter race/ethnic subpopulations through birth or immigration and exit through death or emigration—just as they do in the broader population. Unlike individuals in the total population, however, they can also transition in or out of racial groups by switching identities, a process we refer to as *interracial mobility*. Under these circumstances, natural increase and international migration provide a biased portrait of population change to race/ethnic groups.[2] It is therefore necessary to include the error term $\varepsilon_k$, which restores the equality between the right and left sides of the balancing equation. This term, referred to as the *error of closure* (EOC) in demographic accounting (Preston et al. 2001), can arise for two reasons: measurement error and model misspecification.

The former occurs whenever data from various parts of the balancing equation are drawn from different sources. One popular approach, for example, is to compare the observed population change between censuses with the expected changes based on vital statistics and other independent data sources. This comparison is the foundation of a highly influential method of population accounting known as demographic analysis (Coale 1955), which remains a popular tool for census coverage evaluation to this day (Himes and Clogg 1992). Like the census, demographic data also contain inaccuracies, and the residual term absorbs these errors as well as errors in census coverage.

The EOC also absorbs any components of population change that are not specified in the balancing equation, which defines total population change as the sum of its components. From this zero-sum property it follows that any input left out of the balancing equation can be approximated by subtracting known sources of change from the total population change in a given time period. By this logic, the EOC provides a reasonable, albeit indirect, estimate of unknown sources of population change, provided known sources are measured with little error. This estimation strategy, known as the *residual method* (hereafter RM), has a rich intellectual tradition among demographers, who have long relied upon indirect estimates of unmeasured sources of population change.

[2] Broader population change would not be biased by persons exchanging identities between racial subpopulations, since the total number of persons remains unchanged.

Hamilton (1966) reports that population researchers have been using residual methods since the early twentieth century. Demographers routinely subtracted the difference between births and deaths (net natural increase) from total population change to obtain residual estimates of internal migration (Hart 1921; Thornthwaite 1934). The scope of undocumented immigration also came to light through residual analysis, given the larger than expected counts of Latino and Asian immigrants in the 1980 Census (Fay et al. 1988). Estimates of undocumented immigration can be obtained by subtracting legally resident foreign nationals from the total foreign-born population and then adjusting the difference to account for the undercount of undocumented immigrants in the census (Passel et al. 2004; Passel and Woodrow 1987). Current estimates of unauthorized immigration rely on RM estimation of this type (U.S. Department of Homeland Security 2005), as do most estimates of foreign- and native-born emigration (Ahmed and Robinson 1994; Kraly 1998; Mulder et al. 2002; Mulder 2003; Warren and Passel 1987; Warren and Peck 1980).

A handful of studies have also used residual methods to examine unexpected changes in the size of race/ethnic groups. In the early twentieth century, RM was used to estimate rates of "white passing" among persons with mixed European and African ancestry (Hart 1921; Herskovits 1928), and recent studies of American Indians have attributed much of the growth since 1960 to changes in self-classification among persons with partial indigenous ancestry (Eschbach 1993, 1995; Harris 1994; Nagel 1996; Passel 1976, 1996; Passel and Berman 1986; Snipp 1989). A recent report echoes the risks posed by inconsistencies in racial classification more broadly:

> DA [demographic analysis] estimates of net undercount will be biased if persons who are classified as Black in DA are reported as another race in the census. We need to conduct more research to assess the degree of inconsistency and identify ways this "classification error" can be minimized. (Robinson 1996)

While it is important to consider individual preferences, and changes thereto, as a viable source of change to race/ethnic subpopulations, neither the census nor demographic data contain direct measures of identity preference, life histories of race/ethnic identification, or repeated measures of race/ethnicity at the individual level. Since the probability of switching races cannot be calculated without such measures, residual methods cannot be used to estimate interracial mobility directly. Only aggregate shifts in the size and composition of race/ethnic groups are observed. Interracial mobility can be inferred from these shifts, but only after careful evaluation of other potential explanations, including sources of measurement error. Furthermore, aggregate estimates of net interracial mobility must be interpretable in terms of changing preferences for particular race/ethnic identities.

## 3. A REDUCED-FORM POPULATION ACCOUNTING MODEL

Empirically, it is difficult to disentangle identity shifts from errors in the measurement and estimation of population dynamics, since all are absorbed by a common residual term. To minimize competing sources of closure error, we propose a reduced-form equation that estimates residual change among the native born, which we define as the gap between observed and expected survivorship between one census (time 1) and the next (time 2).[3] After adjusting the EOC for intercensal changes in racial measurement and coverage, the remaining closure error provides a reasonable estimate of net interracial mobility for each race/ethnic group.

In formal demographic accounting, each population process is estimated from distinct sources (e.g., birth and death records). Since our goal is to minimize sources of measurement error, we substitute demographic estimates of fertility and migration with observed census counts of each. Changes in the foreign-born population between $t_1$ and $t_2$ provide a net measure of international migration for each race/ethnic group,[4] as shown in

$$M_{kT} = P_{kt_2}^{fb} - P_{kt_1}^{fb}, \tag{3}$$

[3] For the Census 2000 cohort, $t_2$ estimates are derived from the 2006 American Community Survey.

[4] This simplification ignores native-born emigration and pools foreign-born emigration and deaths.

where $M$ represents net migration, $T$ is the length of the interval between $t_1$ and $t_2$, and $P^{fb}$ is the population of foreign-born persons. Similarly, intercensal births to each race/ethnic subpopulation are defined as

$$B_{kT} = \sum_{x=0}^{T-1} P_{kxt_2}^{nb}, \tag{4}$$

where $x$ is age in years and $P^{nb}$ is the population of native-born persons.

By assuming $P_{kt_2}^{fb} - P_{kt_1}^{fb}$ to be equal to $I_{kT} - E_{kT}$, it follows from (2), (3), and (4) that

$$\delta P_{kT} - B_{kT} - M_{kT} = -D_{kT}^{nb} + \varepsilon_{kT}, \tag{5}$$

where $D_{kT}^{nb}$ represents the native-born deaths to each race/ethnic subpopulation (foreign-born deaths are absorbed by $M_k$). In other words, after subtracting births and changes to the foreign-born population between censuses, the balance of population change accrues to native-born mortality and the EOC. Since the three left-hand values in (5) are measured directly from the census, $-D_{kT}^{nb} + \varepsilon_{kT}$ is identified, but deaths are confounded with errors in this combined quantity. Separating the two requires an independent measure of mortality. Our estimates use life table survival ratios to project each race/ethnic population forward through time (Arias 2002, 2006; National Center for Health Statistics 1985, 1997).[5] Since life tables are available for detailed age groups, the total number of deaths is simply the sum of the age-specific deaths to each native-born census population, or

$$\begin{aligned} D_{kT}^{nb} &= \sum_{x=0}^{90+} D_{kTx}^{nb} = \sum_{x=0}^{89} {}_Tq_x P_x^{nb} + {}_Tq_{90+} P_{90+}^{nb} \\ &= \sum_{x=0}^{89} \left[ \left(1 - \frac{L_{x+T}}{L_x}\right) P_x^{nb} \right] + \left[ \left(1 - \frac{T_{90+T}}{T_{90}}\right) P_{90+}^{nb} \right], \end{aligned} \tag{6}$$

where $T$ is equal to the length of the interval between censuses and $x$ equals age in years (subscript $k$ for race/ethnic group omitted from the

[5] Specific survival rates are available only for whites and blacks; population averages are used for other groups.

right-hand side for simplicity). The quantity ${}_Tq_x$ is the probability of dying between the ages of $x$ and $x + T$, which can be interpreted as the proportion of each census-cohort (by age and race) that is not expected to survive to the next time point (see Appendix B for additional details).

Combining observed counts of intercensal fertility and migration with demographic estimates of mortality results in a reduced form estimate of the *expected subpopulation change* between censuses, or

$$E(\delta P_{kT}) = B_{kT} - D^{nb}_{kT} + M_{kT} = \delta P_{kT} - \varepsilon_{kT}. \tag{7}$$

With mortality uniquely identified, so too is the error term, and though interracial mobility is not the only explanation for the EOC, there are fewer competing explanations than would be the case in more comprehensive demographic models in which the residual absorbs multiple sources of measurement error. In fact, of the three major components of population change, only the most accurately measured—mortality—is estimated from demographic sources in our model (Robinson et al. 1993). Using the net change in the foreign-born population eliminates the need to estimate immigration, return migration, and foreign-born deaths from separate demographic data sources. Similarly, our estimate of intercensal fertility circumvents the problem of quantifying potential identity changes among persons who were not alive at both time points.

This approach is not without limitations. The most significant is that identity change among immigrants and children born between censuses are not measured (though the latter would be included in subsequent censuses). Our specification also ignores U.S.-born expatriates, though native-born emigration is fairly trivial at the population level, recent estimates placing the annual exodus at just 18,000 persons (Gibbs et al. 2003).

This minor loss of generalizability must be balanced against the dramatic reduction in the number of competing explanations for the error of closure. By limiting the analysis of residual growth to native-born persons who are alive at both time points, any biases from unmeasured or poorly measured components of fertility and international migration (illegal immigration, underregistration of births, etc.) are eliminated. This formulation forces intercensal agreement between the expected and enumerated populations of the foreign born and young children, so neither group contributes to the EOC. While this simplification reduces our residual analysis to native-born census cohorts, this subset comprises a large segment of the overall population. Even at $t_2$ (1990, 2000, and 2006), an average of 78% of the population is comprised of native-born persons from the previous census (1980, 1990, and 2000, respectively). Restricting our focus to this universe helps ensure that estimates of interracial mobility are conservative and less likely to be conflated with the leading sources of error in intercensal accounting—fertility and international migration (Robinson et al. 1993).

After fixing two of the three major sources of population change, and estimating the third, the balance of intercensal change to each race/ethnic subpopulation falls on the residual term. Since the latter is no longer subject to errors in the measurement of international migration or fertility, and since native-born emigration is assumed rare enough to be safely ignored, only two sources of error remain confounded with estimates of net interracial mobility: mortality and census coverage.

The former is of little concern, since death records are considered among the most complete and accurate of demographic data sources (Robinson et al. 1993:1064). There is evidence that coroners sometimes misclassify race on death certificates, but these errors should have little impact on the overall growth of race/ethnic subpopulations. First, mortality has been in steady decline for decades, and as a proportion of total population change, mortality trails fertility for every race/ethnic group, and it is a distant third to immigration for the fastest growing groups (Asians and Hispanics). Second, while coroner misclassification may bias mortality rates by 10%–20% for small subpopulations (e.g., American Indians and Pacific Islanders), the impact on larger groups is much smaller—1%–3% for Hispanics, blacks, and whites (Rosenberg et al. 1999). Third, even the well-known, and much larger, differences in relative mortality (e.g., black versus white) are dwarfed by the absolute mortality declines that all groups have experienced in recent decades (Arias 2002, 2006). If lingering racial differences in death rates do not affect the growth of race/ethnic populations to a large degree, minor inaccuracies in the measurement of those differences should have only modest effects. This is fortunate, since life tables for non-black minorities are not available (Arias 2006).[6]

[6] Life table data are available only for the total, black, and white populations.

Differential coverage of censuses, however, poses a much more serious challenge to our inference of interracial mobility. Indeed, the error of closure has traditionally been interpreted as a proxy for coverage error since the invention of modern demographic analysis (Coale 1955). Using independent demographic sources, Coale calculated an expected population for 1950 by summing the 1940 population with the projected gains through natural increase and migration over the subsequent decade.[7] After comparing this expected population to the official census count for 1950, Coale found that the latter was nearly 3% smaller than expected, which he attributed to omissions in the census itself.

While demographic analysis is primarily used to estimate census undercount (Himes and Clogg 1992), Coale notes that for race/ethnic subgroups, undercount is not the only explanation for the error of closure:

> The large apparent deficits among nonwhites pose an interesting question: are these deficits the result of omissions of nonwhites or misclassification of some sort? In this instance "misclassification" is perhaps ... inconsistent classification of the same person as between one census and another, as between a birth certificate and a census form, or as between the entry made by a Selective Service Board and by a census enumerator. (Coale 1955:45)

Of the five conclusions reached in Coale's seminal paper, none was discussed at greater length than the potential contribution of identity shifts to the EOCs of racial groups (Coale 1955:48). The magnitudes of these shifts were not identified, however, and even if they had been, any shifts between race/ethnic subpopulations must, by definition, sum to zero for the total population.[8] As a result, changes in racial classification

[7] This summary oversimplifies Coale's work, which drew upon a number of sources (including vital statistics, draft registrations, and immigration data) and projection methods to produce a range of estimates of intercensal change and census coverage.

[8] Blacks may identify as white, or vice versa, but while these shifts may change the number of blacks and whites, the total number of persons remains the same.

would have no bearing on the total population undercount, for which Coale's discovery is best remembered (Anderson and Fienberg 2001).

While coverage error is conflated with interracial mobility in our model, limited estimates of census undercount are available. Primary sources include demographic analysis and postenumeration surveys conducted by the Census Bureau. Results from the two sources are generally comparable, and in the most recent census, coverage improved to the point that various counting errors (omissions, duplications, etc.) largely offset one another, producing net estimates of coverage error that were close to zero in both DA (0.12%) and postenumeration survey (0.49%) estimates (U.S. Census Bureau 2002).

Despite the sophistication and convergence of coverage evaluation methods, debates about whether to "correct" the census for undercount remain heated. Arguments for and against adjustment are often rooted in divisive political and methodological differences about how to reconcile modest but reliable estimates of national undercount with larger but less precise estimates of undercount for minorities, young men, and urban areas (see Anderson and Fienberg 2001 and Skerry 2000 for contrasting views). Though the RM approach described in this paper can be applied to any level of geography for which data are available, we limit our illustrations of interracial mobility to the national level. Unlike the case with smaller levels of geography, the use of adjusted counts at the national level is "neither difficult nor controversial" (Ericksen et al. 1989:927).

Adjusting the EOC for coverage error eliminates the primary confounder of interracial mobility in our reduced-form accounting model. Disaggregating the error term and solving for the unknown portion shows that

$$\varepsilon_{kT} = r_{kT} + c_{kT} = \delta P_{kT} - E(\delta P_{kT}),$$

where $c_k$ and $r_k$ correspond to coverage error and race reporting shifts, respectively, and

$$r_{kT} = \delta P_{kT} - \left(B_{kT} - D_{kT}^{nb} + M_{kT} + c_{kT}\right), \quad \text{where} \quad \sum_{k=1}^{n} r_{kT} = r_T = 0. \tag{8}$$

In other words, net interracial mobility is simply the balance of subpopulation change after international migration, natural increase, and coverage error are subtracted out. The equation also includes the logical constraint that any net flows between race/ethnic groups cancel out at the population level (see footnote 2).

## 4. INTERRACIAL MOBILITY IN THE PAST QUARTER CENTURY: CENSUS AND ACS EVIDENCE

To illustrate the utility of this reduced-form accounting model, we review the growth of major race and ethnic subpopulations in the United States between 1980 and 2006, and present net estimates of interracial mobility among the native born. Data include the .01 PUMS file from the 1980, 1990, and 2000 censuses, as well as the 2006 American Community Survey, the recently phased in replacement for the decennial census long form.[9] All samples are drawn from the IPUMS data archive at the Minnesota Population Center.[10]

Since changes in race/ethnic categorization and measurement may affect the way individuals identify on the census, several steps are taken to facilitate the comparison of race/ethnicity data across years (see Appendix A for questionnaire wording and categories). The most important change occurred on the 2000 census, which allowed respondents to indicate multiple races by "checking all that apply" (Office of Management and Budget 1997). Previous censuses limited respondents to just one race, so counts from earlier years are not directly comparable with data from 2000 and beyond. To address this issue, we use historically compatible race measures that impute single race responses for those who identify multiple races in Census 2000 and the 2006 ACS. These "bridged" race estimates, originally developed by the National Center for Health Statistics (Ingram et al. 2003), and later modified for use with PUMS microdata (Liebler and Halpern-Manners 2008), permit the comparison of contemporary race data with those from earlier censuses.

Census 2000 was also the first to distinguish Asians from Pacific Islanders, who were counted as a single population in 1980 and 1990. Since there is no way to disaggregate the two groups in earlier years,

[9] Race/ethnicity questions on the ACS are identical to Census 2000, and the 2006 ACS was the first to include group quarters.

[10] For more information, see http://usa.ipums.org/usa/.

we revert to the combined "API" (Asian or Pacific Islander) format for all years. The listings of detailed race/ethnicity categories also varied slightly between censuses,[11] though the major categories, to which we limit our analysis, did not change: Federal guidelines designate four major race categories (white, black, American Indian/Alaska Native, and Asian/Pacific Islander) and a separate measure to distinguish Hispanics from non-Hispanics (Office of Management and Budget 1977). We also include the catch-all "some other race" (SOR) category used by the Census, which is a write-in option for respondents who fail to identify a listed category.

"Hispanic" is not formally included as a race but is instead measured in a separate questionnaire item on Hispanic/Latino origin (see Appendix A). Since this configuration vastly overstates the number of "other race" respondents in the census (Perez and Hirschman 2009), we use a "combined" race/ethnicity format that pools all Hispanics and then compares this aggregate group to non-Hispanics from the four major race categories (Hirschman et al. 2000). Since the vast majority of Hispanics identify as white or SOR on the race question (48% and 43% in Census 2000, respectively), the sizes of these two groups are significantly reduced when Hispanics are treated as a quasi-racial category.[12]

Table 1 decomposes the growth of major race/ethnic groups over the past quarter century. The top panel contains the number of persons (in thousands) added or lost through each source of population change (plus errors), while the bottom panel expresses the contribution of each component as a percentage of total population change. Results show that the United States has grown increasingly diverse in recent years, primarily through racial differences in immigration and natural increase. Between 1980 and 2006, 40 million Hispanics and APIs were added, and the combined share of these groups rose from 8% to 20% of the U.S. total. Most of this growth can be attributed to large numbers of recent immigrants and their U.S.-born children. Non-Hispanic whites, by contrast, added just 20 million persons during this

[11] For example, Eskimo and Aleut were listed in 1980 and 1990 but not in 2000.

[12] In the latter case, removing Hispanics basically depletes the population, since more than 97% of the 15 million persons who identify as SOR in Census 2000 are Hispanic.

# TABLE 1
## Components of Change to Race/Ethnic Populations in the United States, 1980–2006

| | Total Population Change(1) | | | | | Foreign-Born Change | | | | | Intercensal Births | | | | | Projected Native-Born Deaths | | | | | Error of Closure | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 |
| **Race (3)** | 21,392 | 33,298 | 17,977 | 72,666 | | 6,628 | 11,555 | 6,726 | 24,909 | | 35,135 | 38,198 | 23,783 | 97,116 | | 18,237 | 20,343 | 13,160 | 51,740 | | −2,134 | 3,888 | 628 | 2,382 |
| White | 7,525 | 9,397 | 3,187 | 20,110 | | −424 | 1,318 | 533 | 1,428 | | 24,734 | 23,808 | 13,578 | 62,119 | | 15,383 | 17,083 | 11,013 | 43,480 | | −1,402 | 1,354 | 90 | 42 |
| Black | 2,914 | 5,989 | 2,829 | 11,732 | | 552 | 949 | 720 | 2,221 | | 5,152 | 6,153 | 3,681 | 14,987 | | 2,229 | 2,456 | 1,541 | 6,225 | | −562 | 1,342 | −32 | 749 |
| AIAN | 429 | 254 | −63 | 619 | = | 7 | 0 | −4 | 3 | + | 356 | 366 | 184 | 906 | − | 75 | 100 | 61 | 236 | + | 141 | −11 | −183 | −53 |
| API | 3,380 | 4,050 | 2,643 | 10,074 | | 2,503 | 2,748 | 1,669 | 6,920 | | 861 | 1,215 | 922 | 2,997 | | 74 | 94 | 82 | 250 | | 91 | 182 | 134 | 407 |
| Hispanic (2) | 7,129 | 13,393 | 9,058 | 29,579 | | 4,010 | 6,454 | 3,624 | 14,088 | | 3,969 | 6,555 | 5,330 | 15,855 | | 470 | 605 | 455 | 1,530 | | −380 | 988 | 558 | 1,166 |
| SOR | 15 | 216 | 323 | 554 | | −20 | 87 | 183 | 250 | | 64 | 101 | 87 | 252 | | 6 | 5 | 7 | 18 | | −22 | 33 | 60 | 70 |

| | Total Population Change (4) | | | | | Foreign-Born Change | | | | | Intercensal Births | | | | | Projected Native-Born Deaths | | | | | Error of Closure | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 | | 80–90 | 90–00 | 00–06 | 80–06 |
| **Race (3)** | 100.0% | 100.0% | 100.0% | 100.0% | | 10.7% | 15.6% | 15.2% | 14.1% | | 56.5% | 51.6% | 53.7% | 55.1% | | 29.4% | 27.5% | 29.7% | 29.4% | | 3.4% | 5.3% | 1.4% | 1.4% |
| White | 100.0% | 100.0% | 100.0% | 100.0% | | 1.0% | 3.0% | 2.1% | 1.3% | | 59.0% | 54.7% | 53.8% | 58.0% | | 36.7% | 39.2% | 43.7% | 40.6% | | 3.3% | 3.1% | 0.4% | 0.0% |
| Black | 100.0% | 100.0% | 100.0% | 100.0% | = | 6.5% | 8.7% | 12.1% | 9.2% | + | 60.7% | 56.4% | 61.6% | 62.0% | − | 26.2% | 22.5% | 25.8% | 25.7% | + | 6.6% | 12.3% | 0.5% | 3.1% |
| AIAN | 100.0% | 100.0% | 100.0% | 100.0% | | 1.1% | 0.0% | 0.9% | 0.2% | | 61.5% | 76.5% | 42.7% | 75.6% | | 12.9% | 21.0% | 14.1 % | 19.7% | | 24.4% | 2.4% | 42.3% | 4.4% |
| API | 100.0% | 100.0% | 100.0% | 100.0% | | 70.9% | 64.8% | 59.5% | 65.4% | | 24.4% | 28.7% | 32.8% | 28.3% | | 2.1% | 2.2% | 2.9% | 2.4% | | 2.6% | 4.3% | 4.8% | 3.8% |
| Hispanic (2) | 100.0% | 100.0% | 100.0% | 100.0% | | 45.4% | 44.2% | 36.4% | 43.2% | | 45.0% | 44.9% | 53.5% | 48.6% | | 5.3% | 4.1% | 4.6% | 4.7% | | 4.3% | 6.8% | 5.6% | 3.6% |
| SOR | 100.0% | 100.0% | 100.0% | 100.0% | | 17.8% | 38.6% | 54.3% | 42.3% | | 56.8% | 44.6% | 25.9% | 42.6% | | 5.7% | 2.2% | 2.1% | 3.1% | | 19.7% | 14.6% | 17.7% | 11.9% |

*Notes:* (1) Population in thousands. (2) Hispanics (of all races) are pooled. Racial groups (white, black, etc) limited to non-Hispanics. (3) Multiracial respondents from 2000 and 2006 data reassigned to most probable single-race category. AIAN = American Indian or Alaska Native; API = Asian or Pacific Islander; SOR = Some other race. (4) Percentages use absolute values in both numerator and denominator; all values are thus positive.

*Sources:* IPUMS 1% microdata from the 2006 ACS and 1980, 1990, and 2000 Censuses: U.S. Life Tables from 1980, 1990, 2000, and 2003.

interval, a slower rate of natural increase largely due to their older age structure. Nearly 44 million native-born whites died between 1980 and 2006, compared with just 1.3 million native-born Hispanics and APIs.

In each of the past three decades, most of the changes in race/ethnic composition result from standard demographic forces. The proportion unexplained by natural increase and international migration is fairly small for most groups in most periods, although some figures suggest that other factors are at work. The most telling example lies at the demographic margins of the population, in the small but fast-growing number of persons that fail to report any listed race/ethnicity. Usually, the reported "some other race" population is quite large (more than 15 million in Census 2000), but this is largely an artifact of excluding Hispanicity from the census race question (Perez and Hirschman 2009). Failing to locate their primary identity among the listed race categories, many Hispanics choose "some other race" by default (Logan 2004; Perez 2008).[13]

Purged of its 97% Hispanic majority, the small number of SOR respondents that remain is limited to non-Hispanics who refuse to identify with any listed race and whose write-in responses cannot be recoded (or imputed) to a listed category. This residual population numbered less than a quarter million in 1980 and remained largely unchanged ten years later. Between 1990 and 2000, however, the SOR population nearly doubled, and roughly 323,000 additional SOR persons were added between 2000 and 2006.

It is very unlikely that high rates of natural increase and migration could sustain the 11% annualized increase observed for the SOR population between 1990 and 2006. Changes in identity preference certainly contribute to this trend, as Table 1 shows that between 15% and 20% of the "growth" in the SOR population cannot be explained by conventional demographic forces. Still, the example is almost too convenient, since the SOR "population" is defined solely by a shared disregard for the race/ethnicity categories listed on the census. Persons who provide write-ins (e.g., Irish or Nigerian) that are subsumed by official categories are simply recoded (e.g., to white or black). Only

[13] The "some other race" checkbox is followed by a line where respondents may write in their identity, and many Latinos use this space to reiterate their detailed Hispanic origin (Mexican, Cuban, etc.).

those who provide uncodable responses remain in the SOR category, and those who do so comprise a fairly eclectic mix.[14]

More compelling evidence of potential identity shifts lies with-in the traditional racial groups, some of which also experience rates of change that exceed the limits of natural increase and immigration. The most remarkable case is the American Indian/Alaska Native population. After growing by 30% in the 1980s, the AIAN population grew at a considerably slower pace in the 1990s and appears to have even declined in recent years. In the absence of exogenous population shocks like famine, war, or diaspora, such an erratic pattern of growth/decline should be viewed with skepticism.

For example, while Hispanics and APIs also grew quickly during the 1980s, over 90% of the growth was due to immigration and fertility. AIANs, by contrast, added almost no immigrants and experienced just 356,000 births, yet total intercensal growth somehow exceeded 429,000 during the 1980s. That there is an unexplained surplus of AIAN persons in 1990, even before deaths are subtracted out, implies that the 1980 census cohort *grew* over time, adding 66,000 new AIAN persons during a decade in which that population was projected to lose an even larger number (75,000) to mortality. Summing these two discrepancies results in a net EOC of 141,000, or roughly 24.4% of AIAN growth between 1980 and 1990.

Improvements in mortality and census coverage could not have accounted for this trend. More likely, the excess "growth" in the AIAN population reflects continuing fluidity in the willingness to report Native American identity among persons of partial American Indian ancestry. Compounding this suspicion is the equally unusual deficit in projected AIAN growth during subsequent decades. As shown in Table 1, there is a small but negative EOC for the AIAN population between 1990 and 2000, which widens sharply between 2000 and 2006, a dramatic reversal of the earlier trend. Six years into the current decade, the total AIAN population has apparently declined by more than 60K persons, despite adding three new births for every projected death (180K/60K). Barring mass emigration or a severe undercount of AIAN persons in

[14] These include regional or national responses that are racially ambiguous (e.g., Brazilian or Eurasian) or uninformative multiethnic responses such as "rainbow" or "mixed." A complete listing of Census Race Codes is provided at: http://www.census.gov/acs/www/UseData/sf/Append_G_2005_Code_List.pdf.

the 2006 ACS, the "disappearance" of 180,000 AIAN persons (nearly 1 in 10 AIANs from Census 2000) over a six-year interval likely results from identity shifts, though the bridging methodology (discussed below) may also play a role.

Blacks also seem to have experienced seemingly implausible rates of growth, though only during the 1990s. According to Table 1, the black population grew by 6 million persons, or 21%, between 1990 and 2000. These gains would imply an increase of more than 2% per year—nearly double the annual growth rate from the previous decade, and considerably higher than the pace of the current decade (1.3% through 2006). Though immigration and natural increase account for 87% of this growth, the EOC among native-born blacks stands at 1.34 million in the year 2000. A surplus of this magnitude would be equivalent to eliminating more than half the deaths to native-born blacks in the 1990s. Even if blacks experienced the same mortality as whites during this decade, a surplus of more than half a million persons would remain (results not shown).

While it is clear that nontraditional sources of population change underlie these anomalous findings, even large errors of closure are not sufficient evidence of interracial mobility *per se.* First, EOCs might simply reflect changes in census coverage over time. There is no clear pattern to the residual differences between race/ethnic groups, and all but one group's EOCs change signs from one interval to the next. Furthermore, the sum of the residuals over time (group totals) are quite small; only Hispanics show a surplus larger than 750,000 for the full 26-year period. Population totals, by contrast, which *should* sum to zero, instead fall in the 2–4 million range in the 1980s and 1990s, and are differently signed in each decade. Interracial mobility might well be occurring, but any aggregate identify shifts that are taking place are obscured by the large cyclical changes in error rates from one census to the next.

Second, even if the EOCs for racial groups were valid, these aggregate errors mask potentially large differences in residual growth between age groups, which raises important questions about the interpretability of the results. While some expectations are intuitive—for example, to expect less stable identities among children and adolescents—the same need not be expected for the working age or the elderly.

Figure 1 illustrates this age pattern of closure error in greater detail, by listing the normalized percentage EOC for each race/ethnic

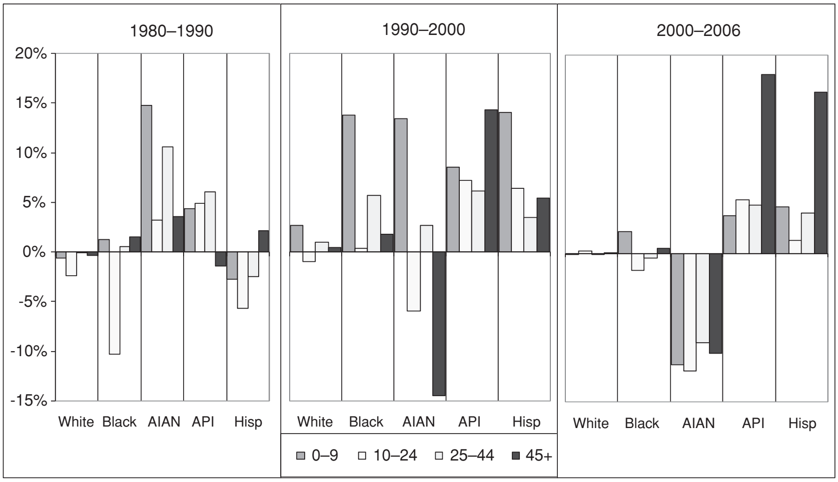


**FIGURE 1.** Normalized percentage EOC by race and age, 1980–2006. Percentages sum to zero across race/ethnic groups.
*Note:* AIAN = American Indian or Alaska Native; API = Asian or Pacific Islander.

group by age.[15] All percentages are weighted for composition and normalized so that summation over $k$ is equal to zero in each interval. While these figures do not account for changes in census coverage, they provide a revealing glimpse of how interracial mobility might play out across different age groups. The most telling pattern is observed for young children. Between 1980 and 1990, for example, survivorship among whites aged 0–9 at $t_1$ fell 1 percent short of expectations at $t_2$, while children from every nonwhite group except Hispanics experienced a residual surplus during the 1980s. For example, the AIAN population aged 0–9 in 1980 was 15% (53,000 persons) larger at $t_2$. This pattern is consistent with a net shift in identification among children who were reported as white in the 1980 census but nonwhite ten years later.

A similar, albeit less pronounced, pattern exists in the most recent interval. From 2000 to 2006, white children again experience a small residual decline, while most nonwhite children (except AIANs) experience residual gains.

[15] Interpreted as the net percent undercount (or overcount) of native-born survivors at $t_2$.

Results for the 1990–2000 interval reveal the largest unexplained gains among nonwhite children in any decade; those who are 0–9 in 1990 grow by 9–14% for every nonwhite group. Even American Indians, who experience a small overall decline during the 1990s (see again Table 1), register a 14% surplus of young children. While these 1990 figures hint at a continuation or acceleration of the pattern set in the 1980s, there is one troublesome distinction: In the '80s, nonwhite gains were logically offset by white losses, but in the 1990s, all children experienced residual growth.

This logical fallacy could be allayed by adding additional constraints (for example, by normalizing the white EOC to zero) or forcing racial differences to sum to zero within age groups, but statistical adjustment will not address the underlying need to disentangle census coverage from interracial mobility in the aggregate EOC. In the 1990s, for example, persons aged 0–9 and 25–44 show net gains for every racial group, and among persons 45 and older, only American Indians appear to decline. All told, just 3 out of 20 race∗age cohorts experience a residual loss in the 1990s. Interracial mobility cannot account for such an erratic pattern. Rather, these results almost certainly reflect the known improvements in census coverage between 1990 and 2000 (U.S. Census Bureau 2002).

To help address this limitation, Figure 2 lists the undercount-adjusted errors for each race/ethnic group in the 1980s, 1990s, and the full interval from 1980 to 2006.[16] From an interracial mobility perspective, the adjusted results are much more intuitive.[17] In the 1990–2000 interval, most nonwhites continue to show a surplus in at least three of four age groups, but whites now show corresponding losses at all ages. The percentages shown in the figure are forced to sum to zero across race/ethnic groups, but after taking changes in census coverage into account, this constraint is almost unnecessary. The unconstrained sum of the group EOCs for 1990–2000 is just 0.69%, less than 250,000 persons (not shown). This is a huge reduction of the unadjusted sum of 3.8 million shown in Table 1. More importantly, just as period totals have grown closer to zero, racial differences across periods have become

[16] Estimates rely on DA adjustment with multiplicative scaling. See Appendix C for details.

[17] Coverage estimates are not available for the American Community Survey, so 2000–2006 values are unadjusted.

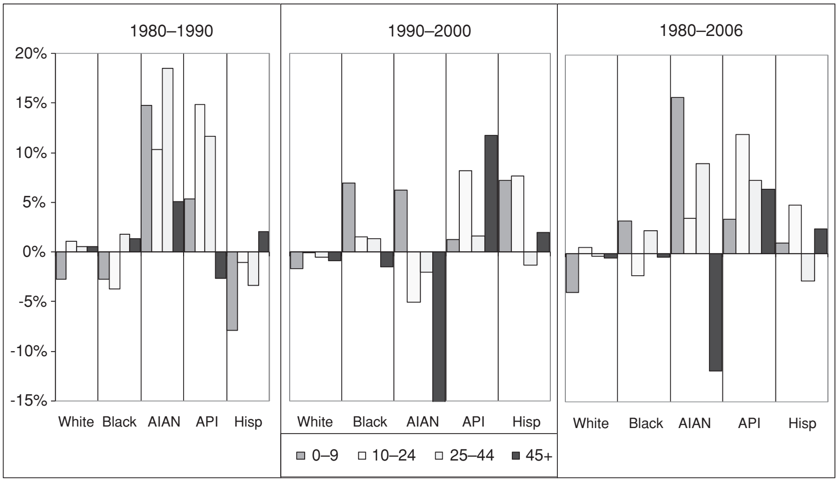


**FIGURE 2.** Net interracial mobility by race and age, 1980–2006. Percentages sum to zero across race/ethnic groups. Coverage estimates unavailable for 2000–2006 (average uses unadjusted values from Figure 1).
*Note:* AIAN = American Indian or Alaska Native; API = Asian or Pacific Islander.

more consistent, which is precisely what should result if identity shifts offset one another within intervals but accumulate over time. Among young children, results for the 1980s are little changed from the unadjusted figures, with the exception of black children switching from a small surplus to a small deficit. For the 1990–2000 interval, the surplus of children among nonwhite groups is now offset by a similar-sized deficit of white children.

Clearly, broad improvements in census coverage were largely responsible for the unadjusted surpluses observed between 1990 and 2000. After adjusting for coverage changes, however, a clearer pattern of net interracial mobility begins to emerge. If the results shown in Figure 2 are accurate, then about half a million children who were identified as white in the 1990 census either identified, or were classified as having, a nonwhite race/ethnicity in Census 2000.

Table 2 puts the magnitude of these reporting shifts into perspective, using WLS estimates of the percentage EOC by race and year. Each observation is a one-year age group from each race and census cohort (N = 486). Model coefficients are the *relative* percent surpluses (or deficits) of persons surviving to $t_2$ for each race, using whites as the

TABLE 2
WLS Estimates of Percentage EOC by Race/Ethnicity, 1980–2006

| | 1980–1990 | | | | | 1990–2000 | | | | | 2000–2006 | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (1) | (2) | (3) | (4) | (5) | (1) | (2) | (3) |
| White (ref) | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Black | −1.5 | −1.3 | 0.1 | 0.1 | 0.1 | 4.5* | 4.4* | 3.4* | 3.4* | 3.4* | −0.2 | −0.1 | 0.0 |
| AIAN | 11.4* | 11.6* | 10.0* | 10.0* | 10.0* | −1.5 | −1.6 | −5.9* | −5.9* | −5.9* | −9.0* | −9.0* | −8.4* |
| API | 7.8* | 7.8* | 5.0* | 5.0* | 5.0* | 7.6* | 6.7* | 9.5* | 9.5* | 9.5* | 3.8* | 3.8* | 6.5* |
| Hispanic | −3.4* | −3.1* | −0.8 | −0.8 | −0.8 | 7.5* | 6.9* | 5.5* | 5.5* | 5.5* | 2.9* | 2.9* | 5.3* |
| Model DF | 5 | 9 | 29 | 30 | 30 | 5 | 9 | 29 | 30 | 30 | 5 | 9 | 29 |
| Model R sq. | 0.11 | 0.22 | 0.34 | 0.35 | 0.35 | 0.14 | 0.26 | 0.36 | 0.37 | 0.37 | 0.11 | 0.12 | 0.18 |

*Note:* Coverage Adjustments derived from Demographic Analysis. AIAN = American Indian or Alaska Native; API = Asian or Pacific Islander. Results for the SOR population included in the estimation, but not reported:

(1) Race-only model, no control variables (Identical to Cross-tab)
(2) Controls for Age Structure (5-Group variable used in earlier tables)
(3) Controls for Racial Differences in Age Structure (Two-way Interaction)
(4) (3) plus control for Intercensal Change in Census Coverage (Additive Scale)
(5) (3) plus control for Intercensal Change in Census Coverage (Multiplicative Scale)
* Percentage shift significantly different than white reference at <0.10
N = 486 (81 age groups by 6 races)
Cells = Weighted Percent EOC minus White EOC (i.e. relative percentage shift)

reference group. All percentages are weighted for composition, and each intercensal panel contains a race-only model (column 1), models that adjust for age structure and racial differences in age structure (columns 2 and 3, respectively),[18] and models that add two different estimates of census undercount.[19] Significant residual shifts are defined as those that differ from the white reference group at the .10 level or better.

Racial differences in residual growth during the 80–90 interval are consistent with earlier tables and figures. Relative to whites, American Indians and Alaska Natives (AIAN) show a net 10–12% residual gain across models, while survivorship among native-born Asians and Pacific Islanders (API) exceeds expectations by about 5%–8%. Furthermore, the negative EOCs (residual losses) for native-born Hispanics are no longer significant after controlling for changes in census coverage. Estimates for blacks are small and insignificant at the .10 level. Summing up the residual shifts during the 1980–1990 interval, results suggest that nonwhite groups either retained their populations or absorbed a small number of non-Hispanic whites during this period.

Results for subsequent years show a more pronounced drift toward minority groups, with the exception of AIANs. APIs and Hispanics experience residual surpluses of 5%–10% after adjusting for differences in age structure (column 3). In the 1990s and 2000s, blacks also experience unexplained growth after 1990, though the surplus is limited to the 1990–2000 interval, and shrinks from 4.5% to 3.4% after controlling for changes in undercount. Table 2 also confirms the rapid fall-off of the double-digit growth in AIAN identities observed during the 1980s. In fact, after adjusting for changes in undercount, the AIAN cohort from the 1990 census appears to have declined by around 6% during the subsequent decade, while the 2000 cohort has lost an even greater percentage (8–9%) as of 2006.

## 5. LIMITATIONS OF THE REDUCED-FORM ACCOUNTING MODEL

These results reveal several instances in which race/ethnic subpopulations change in ways that defy demographic principles and yet cannot

[18] Five category age variable with cutpoints at age 5, 10, 25, and 45.

[19] See Table 2 notes and Appendix C for details on undercount methodology.

be fully explained by census coverage. The overall pattern is tentatively consistent with a net reshuffling of identities, among children in particular, from non-Hispanic whites to various race/ethnic minorities in recent decades. These shifts are modest for certain age groups and time periods, though our estimates of interracial mobility are generally much smaller than the changes accruing to conventional forces of population change (immigration and natural increase).

As proof of principle, the residual approach derived in this paper offers a promising extension to well-known methods of demographic accounting. Substantive interpretations of the empirical results should be drawn with caution, however. First, even after attempting to discount changes in census coverage (a challenging task in itself, as discussed below), closure errors persist for some older age groups. With few *a priori* reasons to expect adults to systematically change their race/ethnicity from one census to the next, the EOCs among older persons might simply be due to measurement error—either in the life table survival rates, estimates of census coverage, or both. Fortunately, patterns observed for young children are more consistent and generally, though not always, of greater magnitude than those observed for adults. More importantly, all but the youngest children in our analysis are enumerated (at $t_2$) after recently entering or passing through adolescence, an age in which race/ethnic identities have been shown to be relatively fluid (Harris and Sim 2002; Perez 2008).

Second, our sample restrictions exclude two important segments of the population (foreign-born persons and children born between censuses) and dismiss a third (native-born emigrants) on the assumption that it is small enough to ignore (Gibbs et al. 2003). Still, while the exclusion of U.S.-born emigrants may inflate estimates of residual growth for all groups, the omission of immigrants and intercensal births likely has an opposite and much larger effect. Given the instability of identities among children and the lack of familiarity with American race/ethnic categories among new immigrants, excluding these groups ignores two potentially large sources of interracial mobility in the United States.

Third, although our reduced-form model draws upon fewer data sources than conventional demographic models, none of the data we utilize were designed to explore patterns of interracial mobility *per se*. The race categories in the life-tables are not directly comparable with those used in the census, and changes in measurement within the census itself threaten the comparability of race across years. We attempt to mitigate this latter concern by using bridged race measures derived from a social survey in which respondents who identify multiple races are asked to choose which single race they would report under the old "check one race only" format (Ingram et al. 2003). The proportion of persons who opt for each single-race response are then used as bridging parameters to link contemporary race data with those from earlier censuses.

Decisions about how to simplify identities depend on a number of factors, however, and the bridging models used to assign single race responses often include rich contextual measures in the prediction equation.[20] Since many data sources, including the PUMS used in this paper, lack detailed geographic measures, bridged race estimates may lack precision (Liebler and Halpern-Manners 2008). Additionally, while the decisions to pool all Hispanics regardless of race(s) and recombine the Asian and Pacific Islander populations reduce the number of cases that require bridging to about 1.4% of the total 2000 and 2006 populations, this solution remains less than optimal for some groups. AIANs, for instance, have a large share (up to one-third) of bridged cases in 2000 and 2006, which suggests that the erratic fluctuations in the residual growth of this population may result, in part, from limitations of the bridging methodology itself.

Fourth, while disparate rates census coverage for minorities and urban areas generate frequent headlines (and controversy) (Skerry 2000), there is little published research that provides detailed breakdowns of census undercount over time. Many sources list aggregate undercounts by race, sex, and age, but few look at age-specific rates *within* race (or race*sex), and detailed rates for nonblack minorities are rarely reported. The authors were unable to locate detailed undercount rates from the 1980 postenumeration program or the 1990 postenumeration survey, or race*age (1-year age groups) estimates from Demographic Analysis for years other than 1990 (Robinson et al. 1993). As a result, the race*age undercount estimates used in this paper are extrapolated from very limited sources (see Appendix C for details). Still, this is more of a data shortcoming than a methodological one. As illustrated in the age-decompositions from Figures 1 and 2, our model can accommodate additional subpopulation characteristics by indexing race/ethnic

[20] Ingram et al. (2003) includes restricted county-level measures of racial groups, as well as the proportion Hispanic and multiracial groups.

groups by other measures of interest (such as age and sex). Racial differences in undercount, for example, are typically more pronounced among males (Anderson and Fienberg 2001). If race*age*sex-specific coverage data were available for multiple census years, gender-specific estimates of interracial mobility could readily be calculated.

Finally, although our reduced-form model attempts to minimize competing explanations for the native-born error of closure, the fairly persistent 5%–8% residual growth observed for native-born Hispanic and Asian adults suggests that additional forms of mobility may be at work. A likely explanation is the misreporting of nativity status. Hamilton (1966) notes that demographic projections are biased by a small number of foreign-born persons who report themselves as native-born on the Census. Given the rapid increases in immigration in recent decades and the often heated public outcries toward undocumented migrants (Lee and Ottati 2002), it is plausible that immigrants may opt to misreport their place of birth on later censuses. Counting these persons as native-born at $t_2$ will inflate native-born survivorship, and by extension, estimates of interracial mobility toward race/ethnic groups with a large share of immigrants (Hispanics and Asians).

## 6. CONCLUDING THOUGHTS

There is considerable speculation about the future race and ethnic composition of the United States. In a recent report, the Census Bureau projects that non-Hispanic whites will fall below 50% of the population by 2042 (U.S. Census Bureau 2008), leading observers in the popular media to claim that a "majority minority" United States is just one generation away (Roberts 2008). While claims about America's racial future often draw upon published U.S. Census reports that are echoed in scholarly writings, it is impossible to predict the race and ethnic composition of the United States without relying upon strong assumptions (Hirschman 2002). Not only are future trends in natural increase and international migration likely to differ from the past and present, but more importantly, race and ethnic identities will remain subjective and socially contextual characteristics. The boundaries between groups are in flux because of intermarriage and the incorporation of new immigrants (Alba and Golden 1986; Farley 1999), and measured changes in racial composition are compounded by structural changes in the measurement and categorization of race in the census and social surveys, as well as selectivity in the assignment of identities to mixed-ancestry children by their parents (Perlmann and Waters 2002; Xie and Goyette 1997). These identities may change again when these children reach adulthood and begin to report their own race/ethnicities (Lieberson and Waters 1993).

Though the future remains uncertain, recent changes in the racial landscape have been swift and pronounced. The most dramatic trend has been the rapid increase in the Asian and Hispanic populations, which grew from just 8% of the population in 1980 to nearly 20% in 2006. The share of African Americans has also inched up during this period, from less than 12% to almost 13% of the American population. Non-Hispanic whites have also increased in number and still comprise two-thirds of the American population, but unlike the immigrant-driven growth of Hispanics and Asians, growth among whites has been largely confined to natural increase, and the gains have been modest because of an older age structure that generates nearly as many deaths as births.

While international migration and natural increase have been the primary forces of change, mobility across race/ethnic categories also appears to have played a role. Race/ethnicity is not simply an ascriptive trait passed down from one generation to the next. America's racial history echoes a complex and as yet unsettled narrative, rife with historical flows across the black-white color line (Davis 1991; Frazier 1937), the absorption of indigenous populations and subsequent resurgence of native identities (Nagel 1996), the declining homogeneity of Asians and Hispanics through intermarriage (Qian and Lichter 2007), and the steady disappearance of white ethnic identities among European-origin groups once thought to be separate "races" (Baltzell 1964; Higham 1988; Waters 1990).

Extending the logic of residual methods used to estimate undocumented migration and other elusive demographic processes, this paper derives a reduced-form population balancing equation that can be used to assess net mobility across race/ethnic groups. Population accounting relies on a formal identity that decomposes sources of population change over time. Any growth left unexplained by measured components—the error of closure—is both a proxy for unmeasured components and the cumulative sum of measurement errors. We attempt to disentangle the two by combining simplifying assumptions with sampling restrictions. After treating migration and fertility as given, net interracial mobility remains confounded with changes in census coverage and the measurement of race/ethnicity. We attempt to account for both factors by combining undercount-adjusted data with historically compatible race measures. The sum of these adjustments yields errors of closure that provide "best guess" estimates of net interracial mobility.

Applying these methods to recent U.S. Census and ACS data, we observe relatively small shifts into race/ethnic minority groups from the non-Hispanic white population between 1980 and 2006. This pattern is most pronounced among young children, half a million of whom may have switched from white to nonwhite in the 1990–2000 interval alone. The major exception to this trend is the American Indian population, which experienced large residual growth from 1980 to 1990, but not afterward. There are at least two explanations for these patterns. One is that some mixed ancestry children are classified by their parents as white when they are young but nonwhite when they reach an age at which they can negotiate their identities with their parents or respond to census and survey inquiries for themselves. The other possibility is that immigrants may claim to be native-born after spending time in the United States. As mobility across ethnic and racial boundaries becomes a larger demographic phenomenon in the years ahead, demographic accounting tools can be adapted to explore these patterns in greater detail.

# APPENDIX A: CENSUS AND ACS RACE/ETHNICITY MEASURES, 1980–2006

## 1980

| | |
|---|---|
| 4. Is this person –<br>*Fill one circle.* | White<br>Black or Negro<br>Japanese<br>Chinese<br>Filipino<br>Korean<br>Vietnamese<br>Indian (Amer.)<br>*Print tribe* →<br>Asian Indian<br>Hawaiian<br>Guamanian<br>Samoan<br>Eskimo<br>Aleut<br>Other—*Specify* |
| 7. Is this person of Spanish/Hispanic Origin or descent?<br>*Fill one circle.* | ○ No ( not Spanish/Hispanic)<br>○ Yes, Mexican, Mexican-Amer., Chicano<br>○ Yes, Puerto Rican<br>○ Yes, Cuban<br>○ Yes, other Spanish/Hispanic |

## 1990

| | |
|---|---|
| 4. Race<br>Fill ONE circle for the race that the person considers himself/herself to be.<br>If Indian (Amer.), print the name of the enrolled or principal tribe. →<br>If Other Asian or Pacific Islander (API), print one group, for example: Hmong, Fijian, Laotian, Thai, Tongan, Pakistani, Cambodian, and so on. →<br>If Other race, print race. → | ○ White<br>○ Black or Negro<br>○ Indian (Amer.) (Print the name of the enrolled or principal tribe.)<br>○ Eskimo<br>○ Aleut<br>Asian or Pacific Islander (API)<br>○ Chinese<br>○ Filipino<br>○ Hawaiian<br>○ Korean<br>○ Vietnamese<br>○ Japanese<br>○ Asian Indian<br>○ Samoan<br>○ Guamanian<br>○ Other API<br>○ Other race (Print race) |
| 7. Is this person of Spanish/Hispanic origin?<br>Fill ONE circle for each person.<br>If Yes, other Spanish/Hispanic, print one group. → | ○ No (not Spanish/Hispanic)<br>○ Yes, Mexican, Mexican-Am., Chicano<br>○ Yes, Puerto Rican<br>○ Yes, Cuban<br>○ Yes, other Spanish/Hispanic<br>(Print one group, for example: Argentinean, Colombian, Dominican, Nicaraguan, Salvadoran, Spaniard, and so on.) |

2000 Census and 2006 American Community Survey

5 Is this person Spanish/Hispanic/Latino? *Mark* ☒ *the "No" box it not Spanish/Hispanic/Latino.*

- ☐ No, not Spanish/Hispanic/Latino
- ☐ Yes, Mexican, Mexican Am., Chicano
- ☐ Yes, Puerto Rican
- ☐ Yes, Cuban
- ☐ Yes, other Spanish/Hispanic/Latino—*Print group.*

6 What is this person's race? *Mark* ☒ *one or more races to indicate what this person considers himself/herself to be.*

- ☐ White
- ☐ Black, African Am., or Negro
- ☐ American Indian or Alaska Native—*Print name of enrolled or principal tribe.*
- ☐ Asian Indian
- ☐ Chinese
- ☐ Filipino
- ☐ Japanese
- ☐ Korean
- ☐ Vietnamese
- ☐ Other Asian—*Print race.*
- ☐ Native Hawaiian
- ☐ Guamanian or Chamorro
- ☐ Samoan
- ☐ Other Pacific Islander – *Print race.*
- ☐ Some other race- *Print race.*

Note: "Negro" not included in 2006 American Community Survey.

## APPENDIX B: MORTALITY PROJECTIONS

As noted in equation (6) in the text, interracial mortality is estimated as

$$
\begin{aligned}
D_{kT}^{nb} &= \sum_{x=0}^{90+} D_{kTx}^{nb} = \sum_{x=0}^{89} {}_{T}q_x P_x^{nb} + {}_{T}q_{90+} P_{90+}^{nb} \\
&= \sum_{x=0}^{89} \left[ \left( 1 - \frac{L_{x+T}}{L_x} \right) P_x^{nb} \right] + \left[ \left( 1 - \frac{T_{90+T}}{T_{90}} \right) P_{90+}^{nb} \right].
\end{aligned}
$$

Note that ${}_{t}q_x$ values are calculated slightly differently for those younger than 90 years of age versus those 90 and older. Among the former, the probability of dying is simply the complement of the forward survival ratio $\frac{L_{x+t}}{L_x}$ for each one-year age-cohort—in other words, the probability that a person of a specific age (e.g., 15) will not survive from one census to the next. Persons older than 89 are top coded at 90+ (the 1980 census maximum) and aged forward using equivalent $T_x$ survival ratios, which are available for persons up to 110 years of age. Since age is top coded in the census, the age structure of the oldest-old is ignored, but the interpretation of intercensal mortality remains the same. In this case, $1 - \frac{T_{90+t}}{T_{90}}$ expresses the probability that all persons aged 90 and older will not survive the interval between measures. For the 1980 and 1990 populations, survival ratios are calculated using the mean of the $L_x$ values from the starting and ending points of each decennial interval (1980–1990 and 1990–2000, respectively), while the 2000 population is projected forward using 2003 life table data, which will have to serve as an approximation of the 2000–2006 average, since more recent data are not available at the time of this writing.

## APPENDIX C: CENSUS COVERAGE ERROR ADJUSTMENT

The quantity $c_k$ is calculated as the difference between the EOCs from the unadjusted and coverage-adjusted counts of each race-age cohort from the 1980, 1990, and 2000 censuses. Table C1 shows undercount estimates by race and year from the two leading sources of coverage evaluation, the Census Post-Enumeration Survey (PES) and Demographic Analysis (DA). While the magnitude and direction of the two

TABLE C1
Percentage Undercount by Race and Year

| | Postenumeration Survey (a) | | Demographic Analysis (b) | | |
|---|---|---|---|---|---|
| Group (1) | 1990 | 2000 | 1980 | 1990 | 2000 |
| White | 0.68 | −1.13 | 0.80 | 1.08 | −0.29 |
| Black | 4.57 | 1.84 | 4.50 | 5.52 | 2.78 |
| AIAN (2) | 5.85 | *0.13* | 4.50 | 5.52 | 2.78 |
| API (2) | *2.36* | *−0.52* | 4.50 | 5.52 | 2.78 |
| Hispanic | 4.99 | 0.71 | 4.50 | 5.52 | 2.78 |

*Source*: (a) U.S. Census Bureau (2003b) and (b) Robinson et al. (1993:1065) and the U.S. Census Bureau (2002:10).

*Notes*: (1) DA available for blacks and non-blacks only. (2) Nonsignificant estimates italicized.

sets of estimates have minor differences, both support the general patterns of (1) higher undercount among minority populations and (2) lower undercount in the 2000 census relative to 1990.

As noted in the text, neither data source contains age-specific estimates of undercount for each race and year, so we extrapolate race × age-specific estimates for 1980 and 2000 using the observed age structure of undercount from the 1990 Census (Robinson et al. 1993), as shown in Table C2. These standardized values are obtained by rescaling the 1990 values to age-specific levels of undercount that, when weighted for composition, reproduce the mean coverage error observed for each racial group in 1980 and 2000. We rescale the 1990 values in two different ways: (1) multiplying each value by a constant, or (2) adding/subtracting a constant from each value. Both transformations preserve the aggregate coverage changes between years (e.g., a net reduction in undercount or overcount). The multiplicative scale also preserves the direction (positive or negative) of coverage errors *for each age group*, as it simply rescales the absolute value of each error to a higher/lower number. By contrast, the additive scale, which adds/subtracts a constant to each age-specific undercount, can change the sign of coverage errors for age groups with observed 1990 values that are close to zero.

The validity of our undercount estimates rests on two potentially problematic assumptions. The first is that undercount rates for blacks and nonblacks can be generalized to minority groups and non-Hispanic

TABLE C2
Percent Undercount by Race, Age, and Year

| | 1980 Estimate Using 1990 Age Structure (3) | | | | Observed Values from 1990 (a) | | 2000 Estimate Using 1990 Age Structure (3) | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Additive Scale | | Multiplicative Scale | | | | Additive Scale | | Multiplicative Scale | |
| Age | Black | Nonblack | Black | Nonblack | Black | Nonblack | Black | Nonblack | Black | |
| 0–4 | 7.51 | 2.36 | 7.08 | 2.07 | 8.28 | 2.54 | 5.53 | 1.22 | 4.17 | |
| 5–9 | 6.71 | 2.36 | 6.39 | 2.07 | 7.48 | 2.54 | 4.73 | 1.22 | 3.76 | |
| 10–14 | 3.16 | 0.21 | 3.36 | 0.31 | 3.93 | 0.39 | 1.18 | −0.93 | 1.98 | |
| 15–19 | −0.79 | −2.39 | −0.02 | −1.81 | −0.02 | −2.22 | −2.77 | −3.53 | −0.01 | |
| 20–24 | 3.21 | −1.09 | 3.40 | −0.75 | 3.98 | −0.92 | 1.23 | −2.23 | 2.00 | |
| 25–29 | 7.91 | 2.91 | 7.42 | 2.52 | 8.68 | 3.09 | 5.93 | 1.77 | 4.37 | |
| 30–34 | 7.86 | 1.86 | 7.38 | 1.66 | 8.63 | 2.04 | 5.88 | 0.72 | 4.34 | |
| 35–39 | 6.16 | 1.01 | 5.92 | 0.97 | 6.93 | 1.19 | 4.18 | −0.13 | 3.49 | |
| 40–44 | 5.16 | −0.04 | 5.07 | 0.11 | 5.93 | 0.14 | 3.18 | −1.18 | 2.98 | |
| 45–49 | 6.26 | 1.26 | 6.01 | 1.17 | 7.03 | 1.44 | 4.28 | 0.12 | 3.54 | |
| 50–54 | 5.91 | 1.21 | 5.71 | 1.13 | 6.68 | 1.39 | 3.93 | 0.07 | 3.36 | |
| 55–59 | 5.16 | 1.31 | 5.07 | 1.21 | 5.93 | 1.49 | 3.18 | 0.17 | 2.98 | |
| 60–64 | 2.81 | 1.41 | 3.06 | 1.30 | 3.58 | 1.59 | 0.83 | 0.27 | 1.80 | |
| 65–69 | −3.69 | 0.26 | −2.50 | 0.36 | −2.92 | 0.44 | −5.67 | −0.88 | −1.47 | |
| 70–74 | −0.74 | 0.51 | 0.02 | 0.56 | 0.03 | 0.69 | −2.72 | −0.63 | 0.01 | |
| 75+ | 2.91 | 0.46 | 3.14 | 0.52 | 3.68 | 0.64 | 0.93 | −0.68 | 1.85 | |
| | 4.500 | 0.800 | 4.500 | 0.800 | 5.520 | 1.080 | 2.780 | −0.290 | 2.780 | Weighted Mean (1) |
| | 0.764 | 0.178 | 0.855 | 0.818 | −0.123 | −0.215 | −2.744 | −1.316 | 0.503 | Scaling Factor (2) |
| | −18.5% | −25.9% | −18.5% | −25.9% | 100.0% | 100.0% | −49.6% | −126.9% | −49.6% | Relative Change from 1990 |

*Source*: (a) Robinson et al. 1993.
(1) Weights = Race*age Composition.
(2) Multiple of (for multiplicative scale) or number subtracted from (for additive scale) observed 1990 values.
(3) 1980 and 2000 estimates are extrapolated from the 1990 race*age structure of undercount.

whites, respectively. The second is that the age structure of undercount did not change appreciably between 1980 and 2000.

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